Research Department Working Paper No:05/01

A Dynamic Model of Central Bank Intervention

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January 2005

The Central Bank of the Republic of Turkey

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Abstract

We examine central bank intervention in foreign exchange markets using a dynamic censored regression model. We allow the amount of purchase and sale interventions to depend nonlinearly upon lagged values of intervention and on measures of disorderly foreign exchange markets. Using data for the CBRT, we find persistence in interventions, which may suggest the presence of political costs and/or a signal of future monetary policy. We find strong evidence of nonnormality and heteroskedasticity in the Tobit model of the reaction function. Results using a robust estimator reveal the importance of considering these specification issues when modeling central bank intervention.

Keywords: Exchange rate, intervention, central bank, dynamic Tobit, CLAD.

JEL classification: F31; F33; G15.

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*We thank the CBRT for providing data on interventions, and especially Emrah Ekşi and Ali Çufadar for their assistance and patience in answering our questions. This paper has benefited from comments by Richard Baillie. The views expressed in this paper do not necessarily represent those of the Central Bank of the Republic of Turkey. All remaining errors are ours.

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1 Introduction

A large body of empirical literature has found evidence that disorderly foreign exchange markets motivate central bank intervention. Yet, because the costs of smoothing out these fluctuations may exceed the benefits, the optimal level of intervention can take on zero values. Therefore, much of the recent literature has focused on specifying and estimating a reaction function that allows for regions of zero intervention in the presence of small variations in the measures of disorderly markets. For example Almekinders and Eijffinger (1994) and Humpage (1999) use a Tobit model to study purchase and sale interventions separately; Almekinders and Eijffinger (1996) and Kim and Sheen (2002) estimate a friction model to explain both types of intervention simultaneously. Alternatively, Baillie and Osterberg (1997), Dominguez (1998), Kim and Sheen (2002); McKenzie (2004), Frenkel and Stadtmann (2001), Frenkel, Pierdzioch and Stadtmann (2003), and Ito and Yabu (2004) estimate discrete choice models for cases when the object of interest is the probability of intervention.

Considerably less attention has been devoted to another aspect of the reaction function, this being the substantial autocorrelation of interventions -purchase (sale) interventions on one day are usually followed by purchase (sale) interventions on the following day. Accounting for this correlation seems to be key to understanding the tactics used by the central bank when intervening in the foreign exchange market. Until not long ago, the difficulty in estimating a dynamic reaction function resided in the lack of formal econometric theory for a model allowing for zero responses of the dependent variable in the face of small changes in the regressors, as well as the inclusion of lagged values of the dependent variable in the latter. Yet, recent work by de Jong and Herrera (2004) establishes the asymptotic correctness of conditional maximum likelihood estimation of the dynamic Tobit model, as well as that of Powell's least absolute deviations (CLAD) estimator for the dynamic censored regression model.

This paper analyzes and estimates the daily reaction function of the Central Bank of the Republic of

Turkey -hereafter CBRT- in deciding the amount of purchases and sales to carry out in the foreign exchange market during the managed float and the free float regimes. This paper is novel in at least three aspects. First, we propose and estimate a reaction function that accounts for the dynamic correlation of interventions and allows for regions of inaction in the face of disorderly markets. We show that dynamics play an important role in the CBRT reaction function, which could be evidence of political costs related to the design of the optimal intervention level and/or the desire of the CBRT to signal its commitment to support the announced exchange rate policy. Second, we test for nonnormality and heteroskedasticity of the disturbances underlying the Tobit model of the reaction function. By using an estimator that is robust to heteroskedasticity and nonnormality -Powell's LAD estimator- we are able to judge the consequences of ignoring these two issues on the estimated reaction function. Finally, estimating the CBRT's reaction function for two subsamples allows us to contrast the motives for intervention under two alternative exchange rate regimes: managed float and free float.

The plan of the remainder of this paper is as follows. In section 2, we describe our intervention data. In section 3, we present our quantitative model for foreign exchange interventions. In section 4, we discuss the econometric techniques. In section 5, we report estimates of the reaction function for the managed float and the free float period. Section 6 concludes.

2 Data

The data set analyzed in this paper contains daily Turkish Lira/US dollar spot exchange rates at 3:30pm, S_t, the interbank overnight interest rate for Turkey, the U.S. federal funds rate, and daily values of foreign exchange interventions by the CBRT separated into purchases and sales. The data on CBRT interventions in the Turkish Lira/US dollar market are measured in millions of US dollars and comprise data for the managed float and the free float regimes. Although, the total data set covers the period between November 1, 1993 and December 31, 2003, we restrict our analysis to two sub-samples well defined by different exchange rate regimes and the absence of financial and exchange rate crisis.[1] These subsamples are: the managed float period between March 1, 1995 and December 31, 1999, and the free float period from February 27, 2001 to December 31, 2003.

Figure 1 plots the development of the daily Turkish Lira/US dollar exchange rate and the daily spot returns ($100 * \Delta \ln S_t$) during the managed float regime, as well as the foreign exchange purchases and sales carried out by the CBRT. The figure shows that during the managed float period the exchange rate experienced a virtually continuous depreciation. Starting at minimum of 39,899 Turkish Lira per US dollar on January 2, 1995, the exchange rate reached a value of 540,098 by December 31, 1999. Spot returns for this period fluctuated between -2.6% and 3.9%, with the largest variations being observed during 1995. Interventions were carried out in 64.5% of the days in the sample with sales (33.9%) being slightly more frequent than purchases (30.6%), particularly during the first two years of the managed float period. Moreover, interventions appear to have been carried out in an effort to smooth movements in the exchange rate. For instance, when the Turkish Lira depreciated at the end of 1995 and then appreciated in January 1996, the CBRT intervened in the foreign exchange market in an effort to dampen the movements in the exchange rate. Note that net interventions (purchases minus sales) took on negative values on the last days of December 1995 and then positive values in January 1996. Moreover, purchase and sale interventions seem to be correlated over time as the CBRT carried out interventions on consecutive days.

The described behavior of interventions appears to be consistent with the exchange rate policy that prevailed during the managed float period. In accordance with the standby agreement signed with the IMF in early 1995, the CBRT's policy was aimed at curbing inflation using the nominal exchange rate as an anchor,

[1] An interesting question that is beyond the scope of this paper, but will be addressed in future reaserch, is unveiling the tactics followed by the CBRT to carry out foreign exchange intervention during financial and exchange crisis.

and particularly at attaining stability in the financial markets. Increases in the foreign exchange basket, defined in terms of German marks and US dollars[2], were supposed to match the target for the monthly inflation rate. Sharp depreciations of the Turkish Lira versus the US dollar, which were followed by large interest rate differentials and increased capital outflows, would have caused pressure on the CBRT to intervene in the foreign exchange market. Furthermore, the inability of the Treasury to achieve a fiscal discipline, combined with the lack of independence of the CBRT, might have constituted an important pressure on the market.

Figure 2 shows the evolution of the daily Turkish Lira/US dollar exchange rate, daily spot returns, and the foreign exchange purchases and sales carried out by the CBRT during the free float period. The figure illustrates how the Turkish Lira/US dollar exchange rate had an upward trend during the first year following the switch in the exchange rate regime. Daily changes in the exchange rate were reflected in the higher variability of the returns during 2001, after the onset of the new regime. In the subsequent years the exchange rate appears to have fluctuated around 1.5 millions Turkish Lira per US dollar, while returns showed a decrease in volatility. The frequency of interventions was lower during the free float period than during the managed float: interventions were carried out only on 50.4% of the days in the sample, with purchases (29.4%) being more frequent than sales (21.0%). Figure 2 provides some evidence that the CBRT carried out interventions in an effort to counter movements in the Turkish Lira/US dollar exchange rate. Interventions carried out during 2001 took mostly the form of sales and were smaller in absolute value than the purchase interventions carried out during 2003. Whereas the frequency of interventions declined during the free float period -in comparison with the managed float period- Figure 2 suggests that interventions continued to be correlated over time.

The beginning of the free float period coincided with the independence of the CBRT and an agreement

[2]The basket was revised to 1 US dollar and 0.77 Euro with the introduction of Euro in 1999.

with the IMF to intervene only in limited amounts in the foreign exchange market. According to IMF (2003), "under the [...] floating exchange rate regime, the Central Bank [was] committed to intervene only to smooth out extreme movements in exchange rates" (2003, p.244), and in agreement with the targets for the foreign exchange reserves. Throughout the free float period, interventions were apparently directed towards dampening excessive volatility in the exchange rate, but without affecting its long-run value. In March 2001 the CBRT began conducting pre-announced (amount and timing) purchase and sale interventions in the form of foreign exchange auctions with the aim of improving the transparency of interventions. Sale auctions carried out in 2001 were intended to sterilize excess liquidity in the market, while purchase auctions in 2002 and 2003 were directed at increasing foreign exchange reserves.[3]

Two common features of foreign exchange intervention are apparent from Tables 1 to 3. First, purchase and sale interventions take on zero values during a large number of business days and a wide range of non-zero values on intervention days. In contrast with interventions in other countries, foreign exchange interventions in Turkey do not appear to be carried out in multiples of a certain amount, but roughly take on a continuum of values.[4] On average, the magnitude of purchase interventions was larger during the managed float period than during the free float period, while the size of sale interventions was smaller.

Second, periods of foreign exchange interventions appear to be followed by periods of intervention, and periods of no activity by similar periods of no intervention. Yet, the magnitude, the frequency, the sign of successive interventions (Table 2) and the average duration between interventions (Table 3) changed between regimes. Note that during the managed float period about 80% of foreign exchange purchases (sales) were followed by interventions of the same sign, whereas during the free float period this probability increased to about 5%. This change across exchange rate regimes is also apparent from the change in the average

[3] See Guimarães and Karacadag (2004) for a detailed description of the different phases of interventions since the floating of the exchange rate.

[4] According to Baillie and Osterberg (1997) the U.S. Federal Reserve carried out interventions in integer multiples of $10 or $100 million of US dollars.

duration between interventions. For the managed float period the average duration between opposite-sign interventions was 3 days, whereas the duration between same-sign intervention was 5 days. The average duration between same-sign interventions remained unchanged during the free float period, while the duration between a first negative and a second positive intervention increased to 18, and that between a positive and a negative intervention was 87 days.

Summarizing, an empirical model of central bank intervention should be able to capture the following features of the data: (i) periods of no intervention in the face of changing conditions in foreign exchange markets; (ii) temporal correlation between interventions: and (iii) a structural change in the reaction function across exchange rate regimes. In the following section we present a model that captures the two first features of the data. As for the changes in regime, we treat the two periods separately.

3 A Quantitative Model of Central Bank Intervention

We use the following dynamic censored regression model to describe the foreign exchange intervention policy of the CBRT:

$$\begin{aligned} INV_t^i &= \max\left(0, \alpha + \sum_{k=1}^{p} \rho_k^i INV_{t-k}^i + \sum_{k=1}^{p} \rho_k^j INV_{t-k}^j + \gamma' x_t + \varepsilon_t\right) \\ &= \max\left(0, \sum_{k=1}^{p} \rho_k^i INV_{t-k}^i + \delta' z_t + \varepsilon_t\right) \end{aligned} \tag{1}$$

where INV_t^i denotes foreign exchange intervention; the superscript $i = B$ $(i = S)$ denotes purchase (sale) interventions; the superscript j denotes the opposite sign intervention, $j = S$ $(j = B)$; α is constant term; x_t is a vector of explanatory variables; and ε_t is a stochastic disturbance. Note that the usual static Tobit model for purchase (sale) interventions is just a restricted version of the dynamic censored regression model

(1) where $\rho_k^i = 0$ for all k, and the ε_t is assumed to be normally distributed.

One reason why this dynamic censored regression model is particularly well suited for analyzing intervention behavior is that it captures the fact that interventions are carried out only on a reduced number of business days. Because the dependent variable remains unchanged in the presence of non-zero values for the regressors for a significant number of business days, estimation of the central bank intervention reaction function cannot be performed using conventional estimators. In fact, OLS estimation would lead to inconsistent estimates. In the past authors have dealt with this econometric issue in different manners. For instance, Almekinders and Eijffinger (1994) and Humpage (1999) estimated two separate Tobit models for purchase and sale interventions; Almekinders and Eijffinger (1996) and Kim and Sheen (2002), among others, used the friction model developed by Rosett(1959) to treat purchase and sale interventions simultaneously. Jun (2004) modified Rosett's friction model in order to capture asymmetries in the response of purchase and sale interventions to disorderly markets, yet in his model the asymmetry depends on the value of one explanatory variable, which acts as a treshold.

Separate Tobit models for purchase and sale interventions allow us to capture two important features of the reaction function: the region of inaction and the asymmetric response of purchase and sale interventions to disorderly markets. However, the usual Tobit model ignores the fact that foreign exchange interventions are correlated over time -once an intervention is carried out, another intervention is likely to take place in the following day. Ito and Yabu (2004) conjecture that this dynamic correlation is due to the presence of political costs associated with the process of designing an optimal intervention policy. More specifically, the cost of intervention on a particular day t may be lower if intervention has been carried out in the previous day. This situation would arise if the central bank has to negotiate interventions with another party and, once an agreement has been reached, interventions can be carried out during several days. This appears to have been the case in Turkey for both periods under analysis.

Temporal correlation can also arise if the objective of the central bank is to minimize an intertemporal loss function that is non-time separable. That would be the case if, for instance, the central bank wants to minimize not only current deviations of the exchange rate from a target, but also past realizations of the deviations. For instance, if the nominal exchange rate is used as an anchor in a deflation program, the central bank would want to reduce the number of times that the target is missed in the same direction in order to maintain its credibility[5]. Consecutive interventions in the foreign exchange market could signal the commitment of the central bank to the announced policy.

Thus, the formulation considered in this paper appears to be appropriate given that 0 values of INV_t^i can capture the optimal level of intervention when the costs of intervention exceed the benefits, and the presence of lagged values of INV_t^i may capture the presence of political costs and/or signalling. Furthermore, the inclusion of lags of opposite sign interventions, INV_t^j, allows us to account for substitutability between purchase and sale interventions.

As for other determinants of foreign exchange interventions, we consider three measures of disorderly foreign exchange markets commonly used in the empirical literature[6]. More specifically, we assume that $\gamma' x_t$ in equation (1) is given by:

$$\gamma' x_t = \gamma_1 \left(s_{t-1} - s_{t-1}^T \right) + \gamma_2 h_{t-1} + \gamma_3 \left(i_{t-1} - i_{t-1}^* \right) \tag{2}$$

where $\gamma' x_t$ is a linear combination of the deviation of the log exchange rate from a log target, $\left(s_{t-1} - s_{t-1}^T \right)$; a measure of excess exchange rate volatility, h_{t-1}; and the interest rate differential between the domestic, i_{t-1}, and the foreign, i_{t-1}^* , overnight interest rates. Note that all the regressors in x_t are lagged one

[5] For a similar argument in the case where the loss function of the central bank is a function of the inflation rate see Bomfin and Rudebusch (2000).

[6] We refer the reader to Sarno and Taylor (2001) for a very completeand recent survey on the topic.

period to avoid simultaneity problems. We now proceed to discuss the motivation and measurement of these regressors.

3.1 Deviations from the target exchange rate

Empirical literature on foreign exchange intervention has found evidence that deviations of the exchange rate from a target Granger-cause intervention. For instance, taking into consideration a very long horizon, Dominguez and Frenkel (1993) show that deviations from purchasing power parity have a significant effect on interventions. When short run considerations -such as a 'leaning against the wind' policy- are taken into account, the target exchange rate is better represented by its past values. This target has been commonly modelled as a moving average of the exchange rate in the past, where the order of the moving average representation varies across studies. For instance, short term horizons of 7, 10 and 25 days have been used by Almekinders and Eijffinger (1996), Humpage (1999), and Frenkel, Pierdzioch and Stadtmann (2003), respectively. Kim and Sheen (2002), Neely (1998), and Le Baron (1999) use a 150-day moving average rule claiming this is a common choice of horizon among market traders. Alternatively, Ito and Yabu (2004) model the target as a weighted average of three past exchange rates in order to capture different horizons.

We model the target as the weighted average of two past representative exchange rates: the Turkish Lira/U.S. Dollar exchange rate in the previous day and its past-month moving average. Thus, deviations from the target are given by:

$$\begin{aligned} s_t - s_t^T &= \gamma_{11}\left(s_t - s_{t-1}\right) + \gamma_{12}\left(s_t - \sum_{j=1}^{20} s_{t-j-1}\right) \\ &= \gamma_{11} dev_t^1 + \gamma_{12} dev_t^{MA}. \end{aligned} \tag{3}$$

The moving average target level for the exchange rate can be thought of as representing past levels of the exchange rate, which is not to say that the latter is considered to be at a desirable level in the previous month. It merely enables us to test whether the central bank systematically 'leaned against the wind' and tried to smooth deviations from the past-month moving average. In addition, it allows interventions to be motivated by daily fluctuations in the exchange rate.

There are two reasons why we do not consider moving averages with longer horizons. First, during the managed float period the CBRT never announced that it would intervene to attain a long-term target. As we mention before, the economic environment and the lack of independence of the CBRT during this period made it more exposed to economic and political pressures to intervene whenever the Turkish Lira depreciated. Once the free float regime was adopted and the agreement with the IMF was in place, interventions where explicitly intended to control short-term volatility in the exchange rate without affecting the long-term exchange rate and taking into consideration the foreign exchange reserve target. Second, our focus on short-term deviations is consistent with results that establish the validity of maximum likelihood estimation for the censored regression model in a time series context (de Jong and Herrera, 2004). In fact, these results hinge on the assumption that the regressors are stationary, which would not be the case for deviations of the exchange rate from medium and long horizon targets.

3.2 Excess exchange rate volatility

Whereas there seems to be ample evidence suggesting that central bank intervention responds to deviations of the exchange rate from a target, evidence regarding the effect of volatility is less conclusive. Baillie and Osterberg (1997) find no significant effect of a GARCH measure of the deviation of conditional volatility from unconditional volatility in the response function of the Federal Reserve Bank and the Bundesbank. Using data for the US, Germany and Japan, Dominguez (1998) finds that exchange rate volatility does not Granger-cause intervention in the 1977-1994 period. On the other hand, Kim and Sheen (2002) find a significant effect of volatility on the interventions of the Reserve Bank of Australia, and Frenkel, Pierdzioch and Stadtmann (2003) reach similar conclusions for the Japanese monetary authority.

Following these studies, we compute a measure of excess exchange rate volatility in the following manner. We assume that the data generating process for the log difference of the exchange rate is given by a GARCH-M(1,1):

$$\Delta s_t = b_0 + b_1 h_t + b_2 INV_{t-2}^{B} + b_3 INV_{t-2}^{S} + b_4 \left(i_{t-2} - i_{t-2}^{*} \right) + \sum_{i=1}^{4} D_{it} \boldsymbol{\lambda}_i + u_t, \tag{4}$$

$$h_t = \omega + \alpha u_{t-1}^2 + \beta h_{t-1} + \gamma_1 INV_{t-2}^{B} + \gamma_2 INV_{t-2}^{S}, \tag{5}$$

$$u_t = \sqrt{h_t} v_t, \tag{6}$$

$$v_t \sim N(0,1), \tag{7}$$

where INV_t^B denotes foreign exchange purchases; INV_t^S denotes foreign exchange sales; $(i_t - i_t^*)$ denotes the overnight interest rate differential; and D_{it} denotes a vector of day of the week dummies. We estimate the model and then measure daily volatility in the foreign exchange market as the estimated conditional variance, h_t, of the daily Turkish Lira/US dollar spot returns. Note that we lag the intervention variables two periods in order to avoid simultaneity.

Table 4 reports estimation results for the GARCH-M(1,1) model (4)-(5), as well as for a GARCH(1,1), corresponding to the managed float period. Regardless of the model specification, the main conclusion that arises from these estimation results is that the CBRT's interventions were not successful in stabilizing the conditional variance or the conditional mean of the Turkish Lira/US dollar exchange rate. This can be seen from the statistically insignificant coefficient on interventions in the conditional variance and the conditional mean equations. Changes in the level of the exchange rate appear to have been driven by interest rate differentials ($t - stat = 2.5$) with increases in the daily exchange rate being on average higher on Mondays

($t - stat = 9.96$) than on other days of the week. We find that ARCH and GARCH effects capture the evolution of the conditional variance and no significant effect of the latter on the conditional mean.

Similarly, during the free float period (see Table 5), foreign exchange interventions carried out by the CBRT had no significant effect on the conditional mean of the Turkish Lira/US dollar exchange rate. For this period, the conditional mean for the exchange rate is well described by a random walk with day-of-the-week effects. As it was the case for the managed float period, the behavior of the conditional variance is well captured by a GARCH(1,1) process. The estimation results reported in Table 5 suggest that foreign exchange sales had a positive but only marginally significant effect on the conditional variance ($t - stat = 1.91$). In other words, sale interventions seem to have increased the volatility of the exchange rate during the free float period. In contrast with the managed float period, interest rate differentials do not appear to have had a statistically significant effect on the behavior of the daily Turkish Lira/U.S. dollar exchange rate. These results are consistent with those of the recent work by Guimarães and Karacadag (2004) who estimate an asymmetric component GARCH model.

One could argue, however, that our finding of a non-statistically significant effect of intervention is driven by the presence of a bias in the maximum likelihood estimates of the conditional mean and conditional variance. Iglesias and Phillips (2004) find that this is the case in small samples when non-zero mean variables (e.g. intervention values) are included in the conditional variance equation of a GARCH model. To make sure that this is not the case, we reestimate our GARCH models using demeaned values of the interventions. Estimation results -not reported herein but available from the authors- are virtually unchanged.

Given our findings, we use as measure of exchange rate volatility the estimated conditional variance of the GARCH(1,1) model for the managed float and the free float period reported in Tables 4 and 5, respectively.

3.3 Interest Rate Differentials

Interest rate differentials constitute another possible force driving central bank intervention. Particularly, the central bank might consider interest rate differentials as a proxy for potential exchange rate overshooting (Kim and Sheen, 2002). Baillie and Osterberg (2000) found that excess-dollar denominated returns over uncovered interest parity (UIP) were associated with purchases of dollar by the Federal Reserve, whereas deviations from UIP had a negative effect on sales of dollars by the German Bundesbank.

In section 3.2 we showed that interest rate differentials had a significant positive effect on the level of the exchange rate during the managed float period. Hence, we would expect that, if the CBRT aimed at smoothing deviations from an exchange rate target, excess returns over UIP would lead to interventions in the foreign exchange market. In this paper, we use the overnight money market rate for Turkey and the U.S. federal funds rate in order to calculate the interest rate differential, $(i_t - i_t^*)$. In this manner we are able to compare interest rates of a 1-day maturity, which avoids the econometric problems related to the use of forward rates with overlapping contracts.

Yet, an alternative measure of Turkey's daily market interest rate calculated by the CBRT is a weighted average of the interest rates in the secondary market for treasury bills and bonds, where the weights are computed according to their transaction volumes. Because Turkey's overnight money market interest rate remained unchanged for large number of consecutive days -particularly during the free float period-, one could argue that this weighted average represent a better proxy for the daily market interest rate. Estimation results using this alternative interest rate not reported in the paper, but available from the authors, are virtually identical.

4 Econometric Techniques

Consider the dynamic Tobit model, that is a censored regression model (1) where the residuals ε_t are assumed to be normally distributed with variance σ^2. Then, the scaled Tobit log likelihood function, $l_T(b)$, conditional on INV_{t-1}^i, ..., INV_p^i, is given by

$$L_T(b) = L_T(\rho, \delta, \sigma) = (T-p)^{-1} \sum_{t=p+1}^{T} l_t(b) \tag{8}$$

where

$$l_t(b) = I(y_t > 0) \log \left(\sigma^{-1} \phi \left(\frac{INV_t^i - \sum_{k=1}^{p} \rho_k^i INV_{t-k}^i - \delta' z_t}{\sigma} \right) \right) + I(y_t = 0) \log \left(\Phi \left(\frac{- \sum_{k=1}^{p} \rho_k^i INV_{t-k}^i - \delta' z_t}{\sigma} \right) \right).$$

De Jong and Herrera (2004) establish that maximizing the log likelihood function (8) over the set of possible parameter values $b \in B$, produces consistent estimates, $\widehat{\beta}_T$, of the dynamic Tobit model. Because $\widehat{\beta}_T$ has an asymptotic standard normal distribution, we can obtain standard errors using the computed Hessian of the log likelihood, or the quasi maximum likelihood estimate of the variance[7]. With regards to the number of lags of the dependent variable, we select the number to be included, p, using Bayes information criterion (BIC).

To obtain estimates of the dynamic censored model that are robust to non-normality and heteroskedasticity we use Powell's CLAD estimator. Powell's CLAD estimator, $\widetilde{\beta}_T$, is defined as the minimizer of the

[7] We report quasi maximum likelihood standard errors.

absolute deviations, $S_T(b)$, over $b \in B$, where

$$S_T(b) = S_T(\rho, \delta) \tag{9}$$
$$= (T-p)^{-1} \sum_{t=p+1}^{T} \left| INV_t^i - \max\left(0, \sum_{k=1}^{p} \rho_k^i INV_{t-k}^i - \delta' z_t\right)\right|.$$

De Jong and Herrera (2004) show that $\widetilde{\beta}_T$ is consistent for the parameters in the censored regression model and has an asymptotic normal distribution with mean β and variance $M^{-1}\Omega M^{-1}$. Because there is no closed form solution to this minimization problem, we compute the CLAD estimates using the iterative linear programming algorithm proposed by Buchinsky (1994). To implement this iterative algorithm, we obtain a first estimate, $\widetilde{\beta}^{(1)} = (\widetilde{\rho}_1^i, ..., \widetilde{\rho}_p^i, \widetilde{\delta}')^{(1)},$ by solving the linear programming (LP) representation of the problem

$$\min_b \left\{ \frac{1}{T-p} \sum_{t=p+1}^{T} \left[\frac{1}{2} sgn\left(y_t - \sum_{k=1}^{p} \rho_k^i INV_{t-k}^i - \delta' z_t\right)\left(y_t - \sum_{k=1}^{p} \rho_k^i INV_{t-k}^i - \delta' z_t\right)\right]\right\} \tag{10}$$

where $sgn(\cdot)$ denotes the sign function. Then, in the second iteration, we restrict the sample to those observations for which the fitted values $\widetilde{\beta}^{(1)\prime} w_t > 0$, where $w_t = [INV_{t-1}^i, ...INV_{t-p}^i, z_t^{'}]'$, and solve the LP problem for $\widetilde{\beta}^{(2)}$ using the restricted sample. We repeat this procedure until the set of observation used in two consecutive iterations is unchanged. Because Powell's CLAD estimator does not provide a first-round estimate of the variance, $M^{-1}\Omega M^{-1}$, we proceed in the following manner. First, we estimate $\widehat{\Omega}$ as the long-run variance of $\psi\left(w_t, \widetilde{\beta}\right) = I\left(\widetilde{\beta}' w_t > 0\right)\left[\frac{1}{2} - I\left(y_t < \widetilde{\beta}' w_t\right)\right] w_t$, where $I(\cdot)$ is the indicator function and $\psi\left(w_t, \widetilde{\beta}\right)$ can be interpreted as a "heuristic" derivative of the objective function $S_T(b)$. We estimate this long-run variance using nonparametric kernel estimation; following the suggestions of Andrews (1991) we use a Bartlett kernel and select the bandwidth according to his formula. To compute $\widehat{M}$, we also obtain an estimate of the conditional density of the innovations in (1) evaluated at zero, $f(0|w_t)$ by nonparametric kernel estimation. In this case we use a higher-order Gaussian kernel estimation with the order and bandwidth selected according to Hansen (2003, 2004).

5 Empirical Results

5.1 Tobit Estimates

Maximum likelihood estimates of the Tobit model, associated standard errors and diagnostic statistics are presented in Table 6 (managed float period) and Table 7 (free float period). We derive three main conclusions: (i) the persistence in interventions suggest that the CBRT had to exert pressure on consecutive days on the foreign exchange market in order to guide the exchange rate in the desired direction; (ii) the response of purchase and sale interventions of the CBRT to disorderly markets was asymmetric; and (iii) the switch in the exchange rate regime was reflected in an increased role of exchange rate volatility in the reaction function.

This empirical evidence underlines the importance of accounting for dynamics in the central bank reaction function. Lags of the dependent variable in both purchase and sale equations are positive and statistically significant. The degree of persistence in interventions decreased in the free float period, as can be seen by the reduction in the magnitude of the coefficients on the lagged dependent variables, and the number of lags selected by the BIC. The dynamic behavior of interventions across subsamples is also reflected in the statistical significance of lags of the opposite sign intervention. Note that during both periods, the coefficients on lagged purchases (sales) have a negative sign in the opposite-sign intervention reaction function. These results may be interpreted as evidence of a signaling channel and/or political costs involved in designing the optimal intervention policy. In fact, the switch to a free float regime and the reduced economic and political pressure associated with the CBRT's independence, are consistent with the decrease in persistence.

As we mentioned above, we derive two main conclusions regarding the effects of "disorderly" markets on foreign exchange intervention: (i) differences between the reaction functions for sale and purchase interventions highlight the importance of accounting for asymmetries; (ii) changes in the exchange rate policy affect the reaction function across subsamples. To substantiate these claims we first compare the reaction functions for purchase and sale interventions for the managed float period. Then we compare the estimation results across exchange rate regimes.

The estimated reaction function for the managed float period (see Table 6) suggest that deviations from the short-run exchange target, dev_{t-1}^{MA}, had a positive and statistically significant effect on foreign exchange sales ($t - stat = 3.02$), whereas the effect on foreign exchange purchases was negative but only marginally significant ($t - stat = -1.83$). These results provide empirical support for the "leaning against the wind" hypothesis of intervention: the CBRT was more likely to carry out sale (purchase) interventions when the Turkish Lira depreciated (appreciated). These results are consistent with the revealed objective of the CBRT -maintaining a stable real exchange rate-, and the inflationary pressure experienced by Turkey during the managed float period. Regarding the conditional volatility of daily exchange rate movements, its effect is negative and statistically significant on foreign exchange sales ($t - stat = -2.51$) and insignificant on purchases ($t - stat = 0.30$). All things equal, the CBRT was more likely to carry out foreign exchange sales when the exchange rate volatility was perceived to be low. Finally, interest rate differentials had a positive effect on purchase interventions ($t - stat = 2.26$), but no statistically significant impact on sales ($t - stat = 1.41$).

Estimation results reported in Table 7 reveal significant changes in the tactics used by the CBRT when intervening in the foreign exchange market during the free float period. First, there is no clear evidence of "leaning against the wind" during the free float period. The coefficient on the change in the exchange rate, dev_{t-2}^{1}, in the purchase (sale) reaction function is positive (negative) and statistically significant. However, this result is reversed when we consider the response of the CBRT's purchase interventions to deviations from the target at a longer horizon, dev_{t-2}^{MA}. Second, for the free float period the effect of the conditional volatility of daily exchange rate movements on purchase interventions is positive and statistically significant ($t-stat = 2.04$). This suggest that, in contrast to the behavior during the managed float period, the CBRT carried out purchase interventions during periods of increased volatility. Regarding interest rate differentials, we find a significant effect on both purchase ($t-stat = -3.15$) and sale ($t-stat = 6.29$) interventions.

Two specification issues that are commonly addressed in Tobit settings are heteroskedasticity and nonnormality (Maddala and Nelson, 1975). The first results in inconsistency of the maximum likelihood estimates of the Tobit model, with the level of censoring being the main determinant of this inconsistency (Greene, 2003). Similarly, nonnormality of the underlying disturbances results in inconsistency of the maximum likelihood estimates. Thus, to test whether the Tobit specification of the reaction function is appropriate, we conduct tests for homoskedasticity and normality.

Tables 6 and 7 report the results for a Lagrange multiplier test of the null hypothesis that the Tobit residuals are homoskedastic (Greene, 2003). This test is obtained by assuming that the residuals exhibit multiplicative heteroskedasticity so that $Var\left(\varepsilon_t|z_t\right) = \sigma^2 \exp\left(\delta' z_t\right)$, where $z_t = \{INV_{t-1}^B, ..., INV_{t-p}^B, INV_{t-1}^S, ..., INV_{t-p}^S, x_t\}$. The test has chi-squared distribution with the degrees of freedom equal to the number of variables in z_t. For the managed float period we reject the null of homoskedasticity at a 1% level with a $\chi_8^2 = 854.2$ (927.3) for purchase (sale) interventions. Similarly, for the free float period, a $\chi_8^2 = 532.0$ (342.2) leads us to reject the null of homoskedasticity for purchase (sale) interventions at a 1% level.

These tables show measures of skewness and kurtosis for the Tobit residuals, as well as the Jarque-Bera statistic. We find strong evidence of non-normality for both subsamples, with nonnormality being more evident for the free float period. Note that with a $\chi_2^2 = 9042.6$ (4032.0) for purchase (sale) interventions, the Jarque-Bera statistic leads us to reject the null of normality for the managed float period at a 1% level. This is also the case for the free float period where the $\chi_2^2 = 127233.7$ (596931.1) for purchase (sale) interventions. The non-normality of the Tobit residuals is also evident in Figure 3. Note that the distributions are highly skewed with a long right tail. In addition, the distribution of the Tobit residuals is leptokurtic relative to the normal. This asymmetry should be kept in mind when comparing the Tobit and the CLAD estimates. In fact, whereas the latter provides estimates that are robust to nonnormality and heteroskedasticity, it is an estimator of the median and not the mean, as it is the case in the Tobit model.

5.2 Powell's CLAD Estimates

Evidence of heteroskedasticity and nonnormality indicate that estimates of the reaction function obtained via maximization of the conditional log likelihood function (8) might be inconsistent. Therefore, we proceed to estimate the dynamic censored regression using Powell's CLAD estimator. Estimation results and associated standard errors are reported in Tables 8 and 9.

Although, the main conclusions regarding the importance of dynamics, the asymmetric response of purchase and sale interventions, and the change in the CBRT tactics, are essentially unaffected, there are two main differences between the CBRT reaction function implied by the Tobit and the CLAD estimates. First, Powell's CLAD produces parameter estimates on the measures of disorderly markets that are larger in absolute value for the managed float period and smaller for the free float period -with the only exception of the parameter on h_{t-1} in the sales equation. These results suggest that violation of the normality and homoskedasticity assumptions in the Tobit model, might lead to understating or overstating the effect of disorderly markets on the CBRT's reaction function.

More specifically, according to the CLAD estimates, interest rate differentials played a significant role in determining the tactics followed by the CBRT when carrying out not only purchase, but also sale interventions. These results point out to a change in the response across exchange rate regimes. While during the managed float period, positive interest rate differentials had a positive and statistically significant effect on purchase ($t - stat = 2.862$) and sale ($t - stat = 4.012$) interventions, during the free float period the effect on sales remained positive and significant ($t - stat = 11.781$), whereas it switched sign on purchase interventions ($t - stat = -3.568$).

Powell's CLAD estimates indicate that both during the managed and the free float regimes, periods of increased volatility in the foreign exchange market lead to heightened purchase and reduced sale interventions by the CBRT. This can be seen from the positive and statistically significant coefficient on h_{t-1} for purchase interventions during the managed and the free float regimes ($t - stat = 2.842$ and 4.297, respectively) and the negative and significant coefficient on sales ($t - stat = -1.849$ and -5.211, respectively). In other words, the CBRT was more likely to carry out purchase interventions when the exchange rate volatility was perceived as being high, and foreign exchange sales when the volatility was perceived as being low.

Second, across both subsamples, the robust estimator suggest a somewhat smaller degree of persistence in foreign exchange interventions. Powell's CLAD estimates on the lagged dependent variables for both periods and interventions are smaller than their Tobit counterparts. Other things equal, the Tobit-estimated average increase in purchase (sale) interventions at time $t + 1$ resulting from a one dollar increase in purchase (sale) of U.S. dollars at time t would be larger than the increase implied the CLAD estimates. This result is consistent with the fact that interventions have an asymmetric distribution with a large right tail (see Tables 5 and 6, and Figure3). In other words, the average value of interventions, which is estimated by the Tobit model, is larger than the median intervention estimated by Powell's CLAD.

As it is the case for the Tobit estimates, the reaction functions implied by the CLAD suggest a significant decrease in the persistence of interventions between the managed and the free float period. Yet, whereas the coefficients on the lags sale interventions in the purchases equation retain their negative signs across estimation methods and exchange rate regimes, this is not the case for all the lags of purchase interventions on the sales equation. Note that whereas the Tobit model clearly imply that purchase interventions constitute a substitute for sale interventions, and vice versa, this is not evident from the CLAD estimates for the free float period. Whereas this might be a result of the bias in the Tobit estimates, an alternative explanation involves possible small sample bias in the CLAD estimates (see Kahn and Powell, 2001). This is an issue that we will be able to address in the future as more observations become available and validity of two-step estimation of censored regression models in a time series context is established.

6 Conclusions

Is central bank intervention effective? In Turkey, foreign exchange interventions appear not to have been effective in altering the exchange rate level, and had only a positive and marginally significant effect on the exchange rate volatility during the free float period. For the managed float period, these results are consistent with high levels of inflation, and the focus of the CBRT's policy on defending the real exchange rate. As for the free float period, our findings are in line with the objective of the CBRT to let the market determine the level of the exchange rate, and only intervene during periods of heightened volatility. Yet, foreign exchange intervention appears to have led to higher, not lower, volatility.

Regarding the motives for intervention, in this paper we show that dynamic considerations play an important role in determining the tactics used by the CBRT. In fact, our estimation results indicate that past values of intervention are informative for predicting current intervention levels. Our findings are consistent with the presence of political costs associated with the design of the optimal intervention policy as suggested by Ito an Yabu (2004), and/or other transaction costs involved in carrying out interventions. Yet, other possible interpretation has to do with the central bank's preferences: persistence in foreign exchange interventions may be related to the desire of the central bank to minimize the number of times that the exchange rate target is missed in the same direction. Indeed, by exerting pressure on the foreign exchange market on a sequence of days the CBRT might have signalled its commitment to the announced monetary and exchange rate policy.

Our results agree with findings of other researchers regarding the importance of excess exchange rate volatility and interest rate differentials in determining central bank intervention. However, the empirical evidence is less conclusive with respect to the role of deviation from an exchange rate target. The magnitude and significance level of these coefficients are particularly sensitive to the estimation method. Differences among the parameter estimates obtained from the Tobit and Powell's CLAD estimator may stem from two sources: (a) inconsistency of the Tobit estimates due to heteroskedasticity and non-normality; (b) a difference between the mean -estimated by the Tobit- and the median -estimated by Powell's CLAD- of interventions, due to an asymmetric distribution of the underlying disturbances. With the caveat that Powell's CLAD estimator in a time series setting might suffer from small sample bias (Kahn and Powell, 2001) -as it is the case in cross-sections-, our findings suggest that future investigations into the motives for central bank interventions should take into consideration the effect of heteroskedasticity and nonnormality on the estimated reaction function.

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Table 1
Foreign Exchange Interventions, Summary Statistics

	All Sample[(a)]		Managed Float[(b)]		Free Float[(c)]	
	Purchases	Sales	Purchases	Sales	Purchases	Sales
Observations	2566	2566	1219	1219	712	712
Mean	85	118	76	106	55	62
Median	40	62	47	69	30	24
Standard Deviation	191	278	88	123	144	72
Maximum	4378	6008	681	879	1517	322
Minimum	0	0	0	0	0	0

[(a)] November 1, 1993 - December 31, 2003
[(b)]January 2, 1995 -December 31, 1999.
[(c)] February 27, 2001- December 31, 2003.

Table 2
Sign of successive foreign exchange interventions

	Sign of second change	
Sign of first change	Positive Intervention	Negative Intervention
Managed Float Period[(b)]		
Positive Intervention	287	73
Negative Intervention	70	238
Free float period[(c)]		
Positive Intervention	151	7
Negative Intervention	9	126

See notes for Table 1.

Table 3
Average Duration Between Interventions

	Sign of second change	
Sign of first change	Positive Intervention	Negative Intervention
Managed Float Period[(b)]		
Positive Intervention	3	5
Negative Intervention	5	3
Free float period[(c)]		
Positive Intervention	3	87
Negative Intervention	18	3

See notes for Table 1.

Table 4
Maximum likelihood estimates of GARCH-in-mean and GARCH models: managed float period$^{(a)}$

	GARCH-M(1,1)		GARCH(1,1)	
	Coefficient	Standard Error	Coefficient	Standard Error
Conditional Mean				
$Constant$	0.009	0.059	-0.003	0.054
h_t	-0.198	0.234		
INV_{t-2}^{B}	-0.0001	0.0001	-0.0001	0.0001
INV_{t-2}^{S}	-7.44E-05	0.0001	-9.23E-05	0.0001
$i_{t-2}-i_{t-2}^{*}$	0.002***	0.0008	0.002***	0.0008
$Monday$	0.329***	0.033	0.328***	0.032
$Tuesday$	0.005	0.025	0.003	0.025
$Wednesday$	-0.004	0.027	-0.004	0.027
$Thursday$	0.009	0.025	0.010	0.026
Conditional Variance				
$Constant$	0.012***	0.003	0.012***	0.003
ε_{t-1}^{2}	0.157***	0.061	0.161***	0.061
h_{t-1}	0.739***	0.048	0.735***	0.046
INV_{t-2}^{B}	-1.78E-06	3.17E-05	-5.13E-06	3.08E-05
INV_{t-2}^{S}	2.74E-05	5.35E-05	2.68E-05	5.41E-05
$Skewness$	0.64		0.62	
$Kurtosis$	9.85		9.55	
Q_{20}	23.83		23.88	
Q_{20}^{2}	13.77		13.84	
$Observations$	1219		1219	

$$\Delta s_t = b_0 + b_1 h_t + b_2 INV_{t-2}^{B} + b_3 INV_{t-2}^{S} + b_4\left(i_{t-2} - i_{t-2}^{*}\right) + \sum_{i=1}^{4} D_{it}\lambda_i + u_t$$

$$h_t = \omega + \alpha\varepsilon_{t-1}^{2} + \beta h_{t-1} + \gamma_1 INV_{t-2}^{B} + \gamma_2 INV_{t-2}^{S}$$

$u_t = \sqrt{h_t}v_t$, with $v_t \sim N(0,1)$

where

$\triangle s_t$ = log difference of daily exchange rate;
INV_t^B = CBRT purchases;
INV_t^S = CBRT sales;
$i_t - i_t^*$ = Overnight interest rate differential;
D_i = Daily dummy (Monday-Thursday);
ε_t^2 = Squared residual;
h_t = Conditional variance of log difference daily exchange rate.

$^{(a)}$Sample: March 1, 1995 through December 31, 1999.

***, **, * denote significance at the 1, 5, and 10% level.

Table 5
Maximum likelihood estimates of GARCH-in-mean and GARCH models: free float period [a]

	GARCH-M(1,1)		GARCH(1,1)	
	Coefficient	Standard Error	Coefficient	Standard Error
Conditional Mean				
$Constant$	-0.141	0.136	-0.175**	0.081
h_t	0.020	0.038		
INV^B_{t-2}	4.06E-05	0.0002	5.87E-05	0.0002
INV^S_{t-2}	0.001	0.003	0.002	0.002
$i_{t-2}-i^*_{t-2}$	-0.001	0.002		
$Monday$	0.200**	0.100	0.200**	0.100
$Tuesday$	0.124	0.098	0.124	0.098
$Wednesday$	0.220**	0.105	0.220**	0.105
$Thursday$	0.081	0.103	0.083	0.103
Conditional Variance				
$Constant$	0.082*	0.045	0.081*	0.045
ε^2_{t-1}	0.245***	0.080	0.244***	0.080
h_{t-1}	0.644***	0.125	0.646***	0.124
INV^B_{t-2}	-4.69E-06	0.0002	-4.30E-06	0.0002
INV^S_{t-2}	0.013*	0.007	0.0134*	0.007
$Skewness$	0.62		0.61	
$Kurtosis$	4.69		4.68	
Q_{20}	26.62		28.05	
Q^2_{20}	15.17		15.29	
$Observartions$	712		712	

$$\Delta s_t = b_0 + b_1 h_t + b_2 INV^B_{t-2} + b_3 INV^S_{t-2} + b_4\left(i_{t-2} - i^*_{t-2}\right) + \sum_{i=1}^{4} D_{it}\lambda_i + u_t$$

$$h_t = \omega + \alpha\varepsilon^2_{t-1} + \beta h_{t-1} + \gamma_1 INV^B_{t-2} + \gamma_2 INV^S_{t-2}$$

$u_t = \sqrt{h_t} v_t$, with $v_t \sim N(0,1)$

where

$\triangle s_t$ = log difference of daily exchange rate;

INV^B_t = CBRT purchases;

INV^S_t = CBRT sales;

$i_t - i^*_t$ = Overnight interest rate differential;

D_i = Daily dummy (Monday-Thursday);

ε^2_t = Squared residual;

h_t = Conditional variance of log difference daily exchange rate.

[a] Sample: February 27, 2001 through December 31, 2003.

***,**,* denote significance at the 1, 5, and 10% level.

Table 6
Tobit estimates of the CBRT's rection function: managed float period[a]

	Purchases of US Dollars (B)		Sales of US Dollars (S)	
	Coefficient	Standard Error	Coefficient	Standard Error
$Constant$	-85.717***	20.929	-163.715***	28.995
INV^B_{t-1}	0.486***	0.069	-0.398***	0.135
INV^B_{t-2}	0.353***	0.064	-0.124	0.128
INV^S_{t-1}	-0.316**	0.145	0.600***	0.085
INV^S_{t-2}	-0.054	0.086	0.311***	0.081
dev^1_{t-1}	-16.128	10.725	13.997	16.08
dev^{MA}_{t-1}	-7.735*	4.237	18.935***	6.266
h_{t-1}	6.042	19.975	-84.464**	33.542
$i_{t-1}-i^*_{t-1}$	0.653**	0.288	0.562	0.401
$LM-Test^{(b)}$	854.227		927.283	
$Skewness$	1.763		0.899	
$Kurtosis$	15.931		11.768	
$Jarque-Bera$	9042.571		4031.982	
$Observations$	1217		1217	

$$INV^i_{t-k}=\max\left(0,\alpha+\sum_{k=1}^{2}\rho^i_k INV^i_{t-k}+\delta' z_t+\varepsilon_t\right),\quad i=B,S$$

$\varepsilon_t|INV^i_{t-1},INV^i_{t-2},z_t\sim N(0,\sigma^2)$

INV^B_t = CBRT purchases;

INV^S_t = CBRT sales;

dev^1_t = exchange rate deviation from previous day rate;

dev^{MA}_t = exchange rate deviation from past-month moving average;

h_t = Conditional variance of log difference daily exchange rate;

$i_t-i^*_t$ = Overnight interest rate differential.

(a) Sample: March 1, 1995 through December 31, 1999.

(b) Lagrange Multiplier Statistic to test for heteroskedasticity.

***,**,* denote significance at the 1, 5, and 10% level.

Table 7
Tobit estimates of the CBRT's reaction function: free float period[a]

	Purchases of US Dollars (B)		Sales of US Dollars (S)	
	Coefficient	Standard Error	Coefficient	Standard Error
$Constant$	73.987**	36.341	-355.652***	50.551
INV_{t-1}^{B}	0.267**	0.128	-2.409***	0.887
INV_{t-1}^{S}	-0.789*	0.410	0.698***	0.128
dev_{t-2}^{1}	23.968***	8.990	-8.425**	3.554
dev_{t-2}^{MA}	-10.334**	4.394	1.558	1.413
h_{t-1}	7.031**	3.452	0.085	1.597
$i_{t-1}-i_{t-1}^{*}$	-4.288***	1.366	5.414***	0.839
$LM-Test^{(b)}$	531.975		342.197	
$Skewness$	5.631		10.283	
$Kurtosis$	68.761		146.042	
$Jarque-Bera$	127233.68		596931.08	
$Observations$	693		693	

$$INV_{t-k}^{i}=\max\left(0,\alpha+\sum_{k=1}^{2}\rho_k^i INV_{t-k}^{i}+\delta' z_t+\varepsilon_t\right),\quad i=B,S$$

$\varepsilon_t|INV_{t-1}^{i},INV_{t-2}^{i},z_t \sim N(0,\sigma^2)$
INV_t^B = CBRT purchases;
INV_t^S = CBRT sales;
dev_t^1 = exchange rate deviation from previous day rate;
dev_t^{MA} = exchange rate deviation from past-month moving average;
h_t = Conditional variance of log difference daily exchange rate;
$i_t-i_t^*$ = Overnight interest rate differential.
[a] Sample: February27, 2001 through December 31, 2003.
[b] Lagrange Multiplier Statistic to test for heteroskedasticity.
***,**,* denote significance at the 1, 5, and 10% level.

Table 8

Powell's LAD estimates of the CBRT's reaction function: managed float period: $^{(a)}$

	Purchases of US Dollars (B)		Sales of US Dollars (S)	
	Coefficient	Standard Error	Coefficient	Standard Error
$Constant$	-38.826***	10.338	-121.864***	28.428
INV_{t-1}^{B}	0.321***	0.018	-6.064***	0.295
INV_{t-2}^{B}	0.146***	0.017	2.666***	0.075
INV_{t-1}^{s}	-0.730*	0.419	0.428***	0.040
INV_{t-2}^{s}	-0.107**	0.048	0.148***	0.038
dev_{t-1}^{1}	-0.161	4.357	14.162	10.895
dev_{t-1}^{MA}	1.755	1.739	-1.755	4.984
h_{t-1}	25.141***	8.846	-61.172*	33.075
$i_{t-1}-i_{t-1}^{*}$	0.410***	0.143	1.695***	0.422
$Skewness$	2.792		3.237	
$Kurtosis$	15.992		22.255	
$Jarque-Bera$	10065.709		20771.125	
$Initial\ observations$	1217		1217	
$Final\ observations$	402		262	

$$INV_{t-k}^{i}=\max\left(0,\alpha+\sum_{k=1}^{2}\rho_k^i INV_{t-k}^{i}+\delta' z_t+\varepsilon_t\right),\ \ i=B,S$$

INV_t^B = CBRT purchases;
INV_t^S = CBRT sales;
dev_t^1 = exchange rate deviation from previous day rate;
dev_t^{MA} = exchange rate deviation from past-month moving average;
h_t = Conditional variance of log difference daily exchange rate;
$i_t-i_t^*$ = Overnight interest rate differential.
$^{(a)}$ Sample: March 1, 1995 through December 31, 1999.
***,**,* denote significance at the 1, 5, and 10% level.

Table 9
Powell's LAD estimates of the CBRT's reaction function: free float period[a]

	Purchases of US Dollars (B)		Sales of US Dollars (S)	
	Coefficient	Standard Error	Coefficient	Standard Error
$Constant$	63.502***	15.560	-37.529***	3.752
INV_{t-1}^{B}	0.192***	0.012	0.233***	0.004
INV_{t-1}^{S}	-0.259**	0.115	0.403***	0.009
dev_{t-2}^{1}	-4.486***	1.211	-1.250***	0.293
dev_{t-2}^{MA}	0.039	0.920	0.735***	0.175
h_{t-1}	3.017***	0.702	-0.745***	0.143
$i_{t-1}-i_{t-1}^{*}$	-1.350***	0.378	0.826***	0.070
$Skewness$	12.300		1.007	
$Kurtosis$	188.530		39.683	
$Jarque-Bera$	1001178.9		38578.237	
$Initial\ observations$	693		693	
$Final\ observations$	221		191	

$$INV_{t-k}^{i} = \max\left(0, \alpha + \sum_{k=1}^{2} \rho_k^i INV_{t-k}^{i} + \delta' z_t + \varepsilon_t\right), \quad i = B, S$$

INV_t^B = CBRT purchases;
INV_t^S = CBRT sales;
dev_t^1 = exchange rate deviation from previous day rate;
dev_t^{MA} = exchange rate deviation from past-month moving average;
h_t = Conditional variance of log difference daily exchange rate;
$i_t - i_t^*$ = Overnight interest rate differential.
[a] Sample: February 27, 2001 through December 31, 2003.
***, **, * denote significance at the 1, 5, and 10% level.

Figure 1: Managed float period

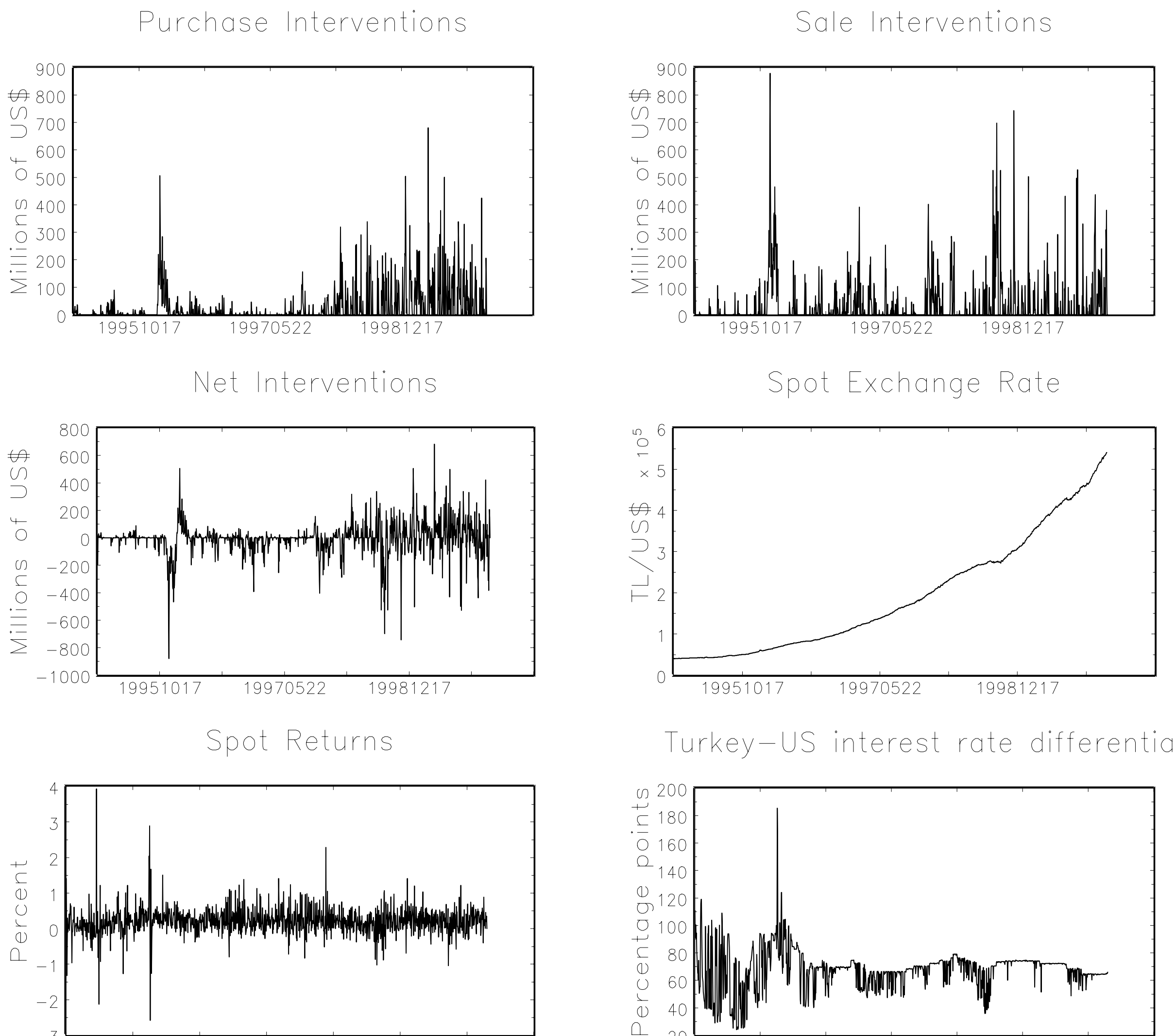

Figure 2: Free float period

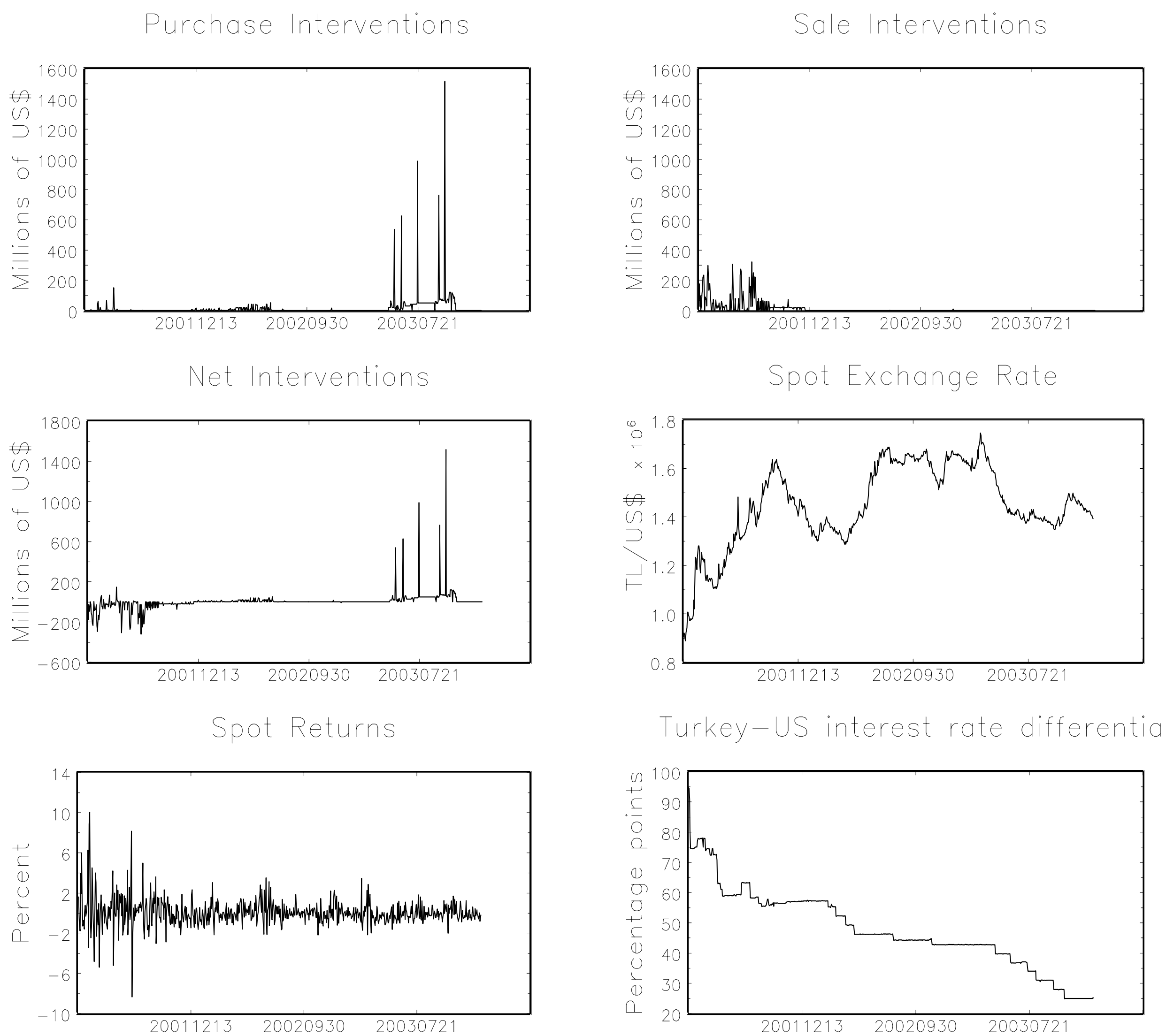

Figure 3: Histogram of Tobit standardized residuals